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van Ommeren, J.N.; Rietveld, P.; Nijkamp, P.

published in

International Regional Science Review
1999

DOI (link to publisher)

[10.1177/016001799761012172](https://doi.org/10.1177/016001799761012172)

document version

Publisher's PDF, also known as Version of record

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citation for published version (APA)

van Ommeren, J. N., Rietveld, P., & Nijkamp, P. (1999). Impacts of Employed Spouses on Job-Moving Behavior. *International Regional Science Review*, 22(1), 54-68. <https://doi.org/10.1177/016001799761012172>

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International Regional Science Review 1999 22: 54

DOI: 10.1177/016001799761012172

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IMPACTS OF EMPLOYED SPOUSES ON JOB-MOVING BEHAVIOR

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This article tests the hypothesis that on-the-job moving behavior differs for the type of household to which the worker belongs. In particular, the authors distinguish between the presence of a spouse and the presence of an employed spouse. They find that female workers with spouses, particularly when they belong to two-earner households, tend to change jobs less often than do other workers. The empirical results do not indicate that job mobility strongly depends on the spouse's workplace location.

The performance of the labor market in terms of vacancies, employment, and unemployment outcomes has been extensively analyzed in numerous studies. Recently, both researchers and policy makers have become increasingly interested in the background factors that influence the degree of flexibility of the labor market. In a flexible labor market, firms can easily fill vacancies and lay off workers, whereas job seekers can easily find jobs that match their skills and tastes. In a flexible labor market, there are fewer obstacles to changing jobs, leading to more labor turnover and, plausibly, less unemployment (Burgess 1992).

In this article, we focus on labor turnover. Labor turnover depends on many structural factors, which differ for each country and change over time (Van Ours 1990). Many studies have focused on the factors that affect unemployed individuals in finding jobs. A smaller number of studies, however, have studied the factors discouraging employed individuals from changing jobs (see, inter alia, Hey and

This research was sponsored by the Economic Research Foundation, which is part of the Netherlands Organization for Scientific Research. We are grateful to Cees Gorter, the editors, and three anonymous referees for useful comments on an earlier version of this article and to the Centraal Bureau voor de Statistiek for providing data. The article was written while the first author was at University College London.



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McKenna 1979; Burgess 1992; Van den Berg 1992, 1995; Hughes and McCormick 1985).

A number of studies have focused on peculiarities in the labor market that may prevent mobility. For example, Hughes and McCormick (1985) provide evidence that occupational pensions reduce labor turnover. Their results, therefore, support the usefulness of the recent legislation of pension transferability designed to increase the flexibility of the labor market in a number of European countries (e.g., Britain and the Netherlands).

Other studies have emphasized that job mobility is affected by factors that reduce residential mobility. For example, Burgess (1992) and Van den Berg (1992) claim that the costs related to moving residences and finding appropriate housing accommodations induce employed persons to change jobs less often. Van den Berg finds that "if one expects it to be hard to sell the present house or to find another house to rent when moving to another job, then job changing costs are (significantly) larger than when such problems are not expected" (p. 1126).

In this article, we intend to investigate whether the job-moving behavior of workers who belong to two-earner households is structurally different from single-wage earners' job-moving behavior. Two-earner households consist of two wage earners who have different working places but share a dwelling, which restricts the choice set of acceptable jobs. One hypothesizes, therefore, that workers belonging to two-earner households change jobs less often than do single-wage earners. Furthermore, this would imply that the spouse's workplace location would affect job mobility (see van Ommeren, Rietveld, and Nijkamp 1998). On the other hand, it may be argued that workers belonging to two-earner households will move more often. For example, the presence of a spouse may reduce the risks involved with changing jobs (e.g., loss of tenure) because the spouse contributes to the household income. This would imply that job mobility does not depend on the spouse's workplace location but on the spouse's wages. Plausibly, these effects are stronger for female than for male workers because male spouses are less likely to leave the labor market. Empirical investigations of the consequences of an employed spouse on job mobility are unknown to us. Previous job mobility studies sometimes include information on the presence of a spouse but generally not the presence of a working spouse. For example, Viscusi (1980) reports that married workers move less often than do single individuals.

The practical importance of emphasizing the presence of employed spouses is evident from the large share of wage earners who belong to two-earner households. For example, according to the Dutch labor force survey, in the Netherlands, about one-third of employed persons are currently part of a two-earner household (EBB 1992), whereas in most other developed countries this share is even significantly higher. This analysis is also relevant for policy purposes. If it is true that two-earner households are less flexible on the housing and labor market, an increase in the share of two-earner households would mean that average commuters will become less sensitive to transportation policy measures aimed at reducing commuting distances.

Although the impacts of employed spouses on job behavior have been ignored in the job mobility literature, these impacts have been examined in the migration literature (see, among others, Sandell 1977; Graves and Linneman 1979; Linneman and Graves 1983; for a more recent contribution, see Ofek and Merrill 1997). It is therefore relevant to make a distinction between migration and other types of job mobility (see Roseman 1971). Migration typically involves a job and residence move over a long distance, so the household leaves the local labor market and living environment. Because almost all job and residential moves are over short distances, the number of moves that may be interpreted as a migration is low.¹ As a consequence, results obtained on the moving behavior of two-earner households based on the analysis of behavior data cannot be generalized to the moving behavior of two-earner households in general. In the context of the effect of an employed spouse on mobility, theories about migration are different from those that deal with all types of moves. A migration of a worker always implies a residential move of the whole household. However, a job move does not necessarily have the same implications.

Although the conclusion obtained by means of an empirical analysis of migration of two-earner households cannot be generalized to the job-moving behavior of two-earner households, in general the results obtained for migration moves indicate that the same result may hold for job-moving behavior. For example, it is well known that single individuals migrate more frequently than married individuals (Mincer 1978). Similarly, the likelihood of a family migration is reduced when the wife of the employee is employed (Sandell 1977). This suggests, but certainly does not imply, that single individuals change jobs more frequently than those who have employed spouses or who are married.

In the job-moving and migration literature, it has been found that certain job characteristics of the spouse have a strong effect on the likelihood of migration and job moving (e.g., the spouse's job tenure). The effect of the spatial location of the spouse's job has been ignored, although there are theoretical reasons to expect that job mobility depends on the spatial location of the spouse's job (van Ommeren, Rietveld, and Nijkamp 1998). As far as we know, this issue has not been empirically investigated.

In summary, in this article we will examine the determinants of job mobility for single-wage earners (with and without spouse) and two-earner households. Our empirical analysis will be based on a hazard model (also called a duration model). We aim to test the hypothesis that on-the-job moving behavior differs for two-earner households and single-wage earners. Moreover, we examine whether job-moving behavior depends on the spatial location of the spouse's job.

The structure of the article is as follows. In section 2, the data and the log likelihood function of the hazard model are discussed. Section 3 contains the empirical results, a discussion of these results, and an examination of the robustness of the results regarding the model specification chosen. Section 4 offers concluding remarks.

THE DATA, THE LOG LIKELIHOOD FUNCTION, AND THE EXPLANATORY VARIABLES

THE DATA

The data set used here (called Telepanel), collected from 1992 to 1993, includes the complete life course pattern of about 3,000 Dutch respondents, including the labor career. The data were collected in a retrospective way. The data set allows for a distinction between voluntary moves and involuntary job moves (due to firing).² From this data set, we have selected 589 persons who worked at least twenty hours per week in the period between 1985 and 1991 and for which all relevant data are observed. A household that consists of two wage earners of whom one works at least thirty-two hours per week (and the other one at least twenty hours) is defined to be a two-earner household.³

One hundred twenty-two observations refer to workers who are part of two-earner households.⁴ We follow the households over time between January 1985 and December 1991 and observe the job durations. After a job move, the household continues to be included in the analysis, so we have multiple job duration observations.

THE LOG LIKELIHOOD FUNCTION

The contributions to the likelihood function are based on multiple duration observations during the observation period of seven years (January 1985 to December 1991). We construct the likelihood function by a repeated stock sampling design on an annual basis. This implies that we sample, at the beginning of each year, the stock of persons who are employed and then observe the labor market transitions of these persons until the end of this particular year. For each year, we observe the elapsed job duration—denoted as p —of the individuals sampled. In case of a completed spell during the year (i.e., a voluntary transition), we observe the spell of residual duration, denoted as r . The total duration is denoted as t . Clearly, $t = p + r$.

We denote the distribution of job duration t as $f(t)$ and its corresponding survival function as $S(t)$, defined as the probability of surviving until t (so, $S = 1 - F$, where F is the cumulative density function [c.d.f.]). We suppose that f and S depend on observed explanatory variables X and an unobserved variable v . Note that $f(t|p)$ can be written as $f(t)/S(p)$.

We use a maximum likelihood method that uses information on the density of the duration t , conditional on the elapsed duration p , $f(t|p)$, a so-called conditional likelihood method. By conditioning on the elapsed duration, maximum likelihood estimation only weakly relies on assumptions about the stationarity of the job-moving process (see Ridder 1984).

In many cases, the completed spell of the job duration is not observed, but it is only known that the duration is longer than a certain value (right censoring). This may happen because during the year in which the worker is observed no transition occurs because another transition is made (e.g., an involuntary move) or because the worker does not belong any more to the same two-earner household (e.g., through divorce or because the spouse becomes unemployed). In the latter case, we will use information on $S(t|p)$, which can be written as $S(t)/S(p)$. So, two types of observations (right censored and completed spells) are included in the likelihood function.

To construct the likelihood function, we integrate over the unobserved variable v using the mixing distribution $h(v)$. The likelihood function L of N ($N = 589$) individuals can then be written as follows ($i = 1, \dots, 7; j = 1, \dots, N$):

$$L = \prod_{j=1}^N \frac{\int \prod_{i=1}^7 [f(t_{ij})^{1 - \text{cen}(ij)} \cdot S(t_{ij})^{\text{cen}(ij)}]^{\text{sam}(ij)} h(v) dv}{\int \prod_{i=1}^7 [S(p_{ij})]^{\text{sam}(ij)} h(v) dv}, \quad (1)$$

where $\text{sam}(ij) = 1$, if individual j is sampled in year i , otherwise 0; $\text{cen}(ij) = 1$, if the spell of individual j in year i is right censored, otherwise 0.⁵

During the period of observation (January 1985 to December 1991), many explanatory variables change. In our specification, we allow the explanatory variables to change annually (the vector X is individual and year specific). The unobserved variable v is individual specific; we treat v as constant over the period under observation.

The distribution $f(t|X, v)$ can be parameterized in different ways (Lancaster 1990). In our empirical application, we employ the mixed proportional hazard model, where the baseline hazard is exponential (and hence, we exclude duration dependence).⁶ This specification implies that

$$f(t|X, v) = \exp^{X\beta} \cdot v \cdot \exp^{-\exp^{X\beta} \cdot v} \quad S(t|X, t) = \exp^{-\exp^{X\beta} \cdot v}, \quad (2)$$

where β is a vector of parameters to be estimated. The hazard rate of changing jobs, denoted as θ , can then be calculated. The hazard rate of changing jobs is the rate of leaving the present job per unit time. Thus, $\theta(t)$ equals $f(t)/S(t)$, which can be written as $\exp(X\beta) \cdot v$.

Estimation of β proceeds generally given an assumption on the form of the distribution of v , $h(v)$, the mixing distribution. We suppose that the mixing distribution $h(v)$ is parameterized with discrete mass points. This means that the heterogeneous sample is endogenously subdivided into homogeneous groups. In the empirical application, we assume that the mixing distribution is such that v has two mass points, and we denote these mass points by v_1 and v_2 . The mixing function consists then of the two probabilities P_1 and P_2 , which are defined as follows: $P_1 = P(v = v_1)$ and $P_2 = P(v = v_2)$. Hence, with probability P_1 , the hazard rate is equal to $\exp(X\beta) \cdot v_1$,

and with probability P_2 , the hazard rate is equal to $\exp(X\beta) \cdot v_2$. Particularly, the flexibility of this distribution is attractive and avoids the need to make strong assumptions on the functional form of the mixing distribution. The statistical model is estimated by a maximum likelihood procedure using the package Gauss.

THE EXPLANATORY VARIABLES

In the empirical analysis, we have used a large number of explanatory variables. We discuss these variables here. The following levels of education are included: university, polytechnic, vocational, high school, and low vocational. Individuals who have only primary school education are in the reference group. Job-to-job mobility is thought to increase with higher educational achievement because higher education offers higher career potential—not only formal education but also the position within the firm affects mobility. Therefore, we include the number of subordinates and whether the person works more than thirty-two hours per week.

We also include the size of the branch. Size of branch is defined as the number of persons working at the same workplace location of the firm. This variable is a proxy for the size of the firm. It is generally thought that larger firms offer more opportunities to grow within the firm and offer better employment conditions, which reduce workers' job mobility. It is generally thought that those who work as civil servants will move less often, so we also include a dummy to capture such an effect.

Calendar year effects are incorporated to capture changes in general labor market conditions by including biannual dummies. In addition, the logarithm of wage rate is included, since a higher wage reduces job mobility (it is common to use the logarithm of the wage instead of the wage in job mobility studies) (see, e.g., Van den Berg 1992, 1995).

Furthermore, we include a range of individual and household explanatory variables:⁷ a dummy for the presence of a spouse and a dummy for whether the spouse is employed (two-earner household), the ratio of the worker's wage over the spouse's wage, a dummy for males, dummies for age groups, and a dummy for the presence of children. Furthermore, the dummy for the presence of a spouse, the dummy for the two-earner household, and the dummy for the presence of children are gender dependent. The time-varying variables are allowed to differ yearly.

We also include information on the commuting distance, z_1 , and the commuting distance of the spouse, z_2 (see Figure 1). We hypothesize that the effect of commuting distance z_1 on job mobility is positive. The effect of the spouse's commuting distance z_2 is ambiguous. (In van Ommeren, Rietveld, and Nijkamp 1998, a simultaneous job and residence search model of two-earner households is introduced. According to this model, the effect of z_2 is ambiguous). Our explanation for this ambiguity is the following: given an increase in z_2 , the household is more likely to move residence. This gives the worker an incentive to accept fewer job offers close to the current residential location and an incentive to accept more job offers far from

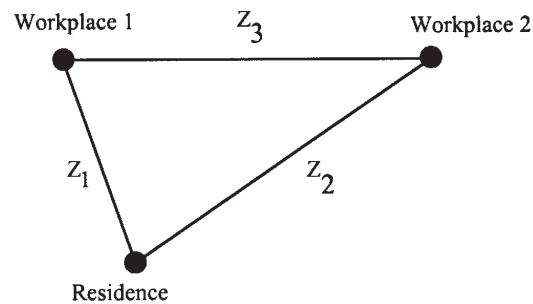


FIGURE 1. The Workplace and Residential Locations of a Two-Earner Household

the current residential location. As a result, the effect of z_2 on job mobility is ambiguous. Nevertheless, we hypothesize that it is likely that the overall effect is positive, since the probability of receiving a job offer at a distance that is longer than the current distance is for most workers larger than the probability of receiving a job offer at a distance that is shorter than the current distance. For example, suppose that employment is uniformly distributed over space: $g(z_i) = c \cdot 2 \cdot \pi \cdot z_i$ for $z_i < z_{\max}$, where the density $g(z_i)$ denotes the distribution of commuting distance offers, z_{\max} is the maximum value of z_i , and c is a normalization constant. Let us consider a worker of whom the current commuting distance is one-fifth of the maximum commuting distance offered.⁸ The probability that a random job offer will reduce the current commuting distance is 0.04. As a consequence, the probability that a random job offer will increase the current commuting distance is 0.96.

In the data set, exact information about the commuting distance is missing, since only data on the municipalities of residence and workplace of the individuals are available. We approximate commuting distance by the distance between the centers of the municipalities (measured in kilometers).⁹ We also include the distance between the workplaces z_3 (see Figure 1). Theoretical reasoning suggests the effect of z_3 on job mobility to be positive (see van Ommeren, Rietveld, and Nijkamp 1998 for a mathematical proof of this reasoning that is based on a simultaneous job and residence search model of two-earner households).

We may explain the effect of z_3 as follows. For smaller values of z_3 , the household is maximally able to reduce the commuting distances of both spouses by means of a residential move. This can be easily understood as follows. Suppose that z_3 is zero, and therefore, both wage earners work at the same location. In this situation, any

residential move that reduces z_1 will also reduce z_2 , which will make a residential move more attractive. As an alternative, suppose that z_3 is equal to the sum of z_1 and z_2 , so the residence location is exactly between the two job locations. In this situation, any residential move that reduces z_1 will simultaneously increase z_2 , which makes moving residences less attractive. So, more generally, the household is better off given smaller values of z_3 , and the household has therefore an incentive to reduce z_3 by means of a job move. Consequently, we hypothesize that for larger values of z_3 , those in two-earner households will change jobs more frequently.

THE EMPIRICAL RESULTS

THE ESTIMATES

The empirical results of the hazard coefficients β can be found in Table 1, model 1.¹⁰ First, we will discuss the variables that are related to the distinction between two-earner households and single-wage earners.

The results appear to indicate that the effect of the presence of a spouse (employed or not) are not significant in general, using a significance level of 5 percent.¹¹ However, a closer look at the results shows that the coefficient for females who belong to two-earner households is significant and equals -1.02 . The latter effect has been calculated as the sum of the estimated coefficients of the dummies for a two-earner, a spouse, two-earner (if female), and spouse (if female). Note that these dummies do not exclude each other. The variance of this estimate equals 0.32 (see the appendix). Thus, female wage earners who belong to two-earner households have significantly lower job mobility than do single-wage earners.

As explained in the introductory paragraphs, a plausible explanation for this effect is that female workers who belong to two-earner households are less flexible in the housing market because they take the workplace location of their spouses into account and thus will not accept job offers from employers at a large distance from their current residences.¹² As a consequence, female members of two-earner households may be less able than single-wage earners to obtain jobs that pay higher wages.

Similarly, it appears that the coefficient for females who have spouses is significant and equals -0.75 (the standard deviation is 0.32 ; for details, see the appendix). Thus, female wage earners with spouses (employed or not) have significantly lower job mobility than do single-wage earners.

A comparison of female workers who belong to two-earner households with male workers who belong to two-earner households indicates that females move less often, although the effect is only significant at the 10 percent level (the effect is -0.70 , and the standard deviation is 0.37 , so the t -test is equal to 1.89).

It is important to notice that, according to our results, on-the-job moving behavior of a female worker does depend on the presence of a spouse, but the effect is

TABLE 1. Empirical Coefficients for Hazard Rates for Leaving Job Voluntarily

<i>Variables^a</i>	<i>Model 1</i> <i>All Observations</i>	<i>Model 2</i> <i>All Observations</i>	<i>Model 3</i> <i>With Spouse</i>	<i>Model 4</i> <i>Two-Earner</i>
Two-earner	-0.12 (0.29)	-0.11 (0.29)	-0.59 (0.28)	
Spouse	-0.20 (0.19)	-0.20 (0.18)		
Two-earner (if female)	-0.15 (0.38)	-0.15 (0.38)	0.15 (0.47)	
Spouse (if female)	-0.55 (0.37)	-0.55 (0.38)		
Children (if female)	0.49 (0.30)	0.48 (0.31)	0.15 (0.47)	0.86 (0.47)
Workplace location of the spouse				
Commuting distance z_2	0.14 (0.06)*	0.11 (0.07)	0.10 (0.10)	0.13 (0.05)*
α (in radians)		0.19	(0.14)	
z_3 (distance in 10 kilometers)	-0.01 (0.09)		-0.02 (0.10)	0.00 (0.12)
Commuting distance z_1 (distance in 10 kilometers)	0.47 (0.24)*	0.46 (0.28)	0.64 (0.25)*	1.88 (0.62)*
Wage rate ^b	-2.19 (0.19)*	-2.23 (0.20)*	-1.05 (0.23)*	-1.26 (0.66)
Wage/wage of spouse	0.01 (0.02)	0.01 (0.02)		
Male	1.41 (0.64)*	1.40 (0.62)*	1.01 (0.60)	0.55 (0.39)
Age				
< age < 24	0.76 (0.29)*	0.76 (0.29)*	1.21 (0.52)*	2.66 (1.06)*
24 < age < 34	0.70 (0.25)*	0.69 (0.25)*	1.11 (0.49)*	1.75 (0.99)
34 < age < 44	0.95 (0.22)*	0.94 (0.22)*	1.10 (0.50)*	1.51 (1.00)
Size of branch				
size > 200 people	-0.75 (0.15)*	-0.76 (0.15)*		-0.48 (0.45)
20 people < size < 200 people	-0.44 (0.12)*	-0.44 (0.12)*		-0.48 (0.37)
More than 32 working hours	-0.13 (0.18)	-0.13 (0.18)		-0.87 (0.47)

TABLE 1 Continued

<i>Variables^a</i>	<i>Model 1</i> <i>All Observations</i>	<i>Model 2</i> <i>All Observations</i>	<i>Model 3</i> <i>With Spouse</i>	<i>Model 4</i> <i>Two-Earner</i>
Number of subordinates				
0	0.13 (0.18)	0.13 (0.18)		-0.42 (0.56)
1, 2, or 3	-0.13 (0.19)	-0.12 (0.19)		0.40 (0.51)
No civil servant	-0.01 (0.06)	-0.01 (0.06)		
Educational level				
University	0.87 (0.27)*	0.88 (0.27)*	0.78 (0.30)*	1.06 (0.53)
Polytechnic	0.08 (0.18)	0.08 (0.18)	0.06 (0.23)	
Vocational	-0.52 (0.19)*	-0.52 (0.19)*	-0.44 (0.19)*	
High school	0.35 (0.24)	0.31 (0.25)	-0.01 (0.25)	
Low vocational	-0.77 (0.23)*	-0.76 (0.23)*	-0.33 (0.25)	
Lives with parents	-0.62 (0.19)*	-0.63 (0.19)*		
Children	0.18 (0.22)	0.18 (0.22)	0.47 (0.34)	0.14 (0.32)
Calendar year				
1985-1986	-0.65 (0.17)*	-0.64 (0.17)*	-0.52 (0.23)*	
1987-1988	-0.36 (0.15)	-0.36 (0.15)	-0.22 (0.23)	
1989-1990	-0.16 (0.15)	0.16 (0.15)	(0.01) (0.15)	
Mass points and probabilities				
v_1	1.37 (1.74)	1.26 (1.64)	0.13 (0.05)*	0.22 (0.61)
v_2	6.99 (1.13)*	5.99 (1.11)*	0.82 (0.10)*	2.27 (0.88)*
P_1	0.99 (0.01)	0.99 (0.01)	0.99 (0.01)	0.99 (0.43)
P_2	0.01 (0.01)	0.01 (0.01)	0.01 (0.63)	0.01 (0.42)
Number of observations	589	589	420	122
Log likelihood	-1,010.61	-1,011.42	-648.16	-119.69

Note: Standard errors in parentheses.

a. Reference groups: type of household (single-wage earner), household situation (no spouse), male (female), age (older than forty-four), size of branch (fewer than twenty), more than thirty-two working hours (less than thirty-two), number of subordinates (more than three), no civil servant (civil servant), educational level (primary and unknown), calendar year (1991).

b. Logarithm of net wage per hour (in Dutch guilders).

* Significant at 5 percent.

stronger given the presence of an employed spouse. Hence, we find that belonging to a two-earner household and the presence of a spouse (employed or not) reduce a female's on-the-job moving behavior. As a consequence, many females may reject job offers that are accepted by male workers. This finding contributes to our understanding of why the wage growth of female workers is less than that of male workers (Sandell 1977; Loprest 1992).

The empirical results provide evidence that the spatial location of the spouse's job affects job-moving behavior. In particular, we find that the spouse's commuting distance z_2 increases job mobility. Such a result is in line with our hypothesis (see the Explanatory Variables section). However, the empirical results do not indicate that the effect of z_3 on job mobility is positive. Consequently, we do not find empirical support for our hypothesis that for larger values of z_3 , those in two-earner households will change jobs more frequently.

The results of the effects of the other explanatory variables in the model are essentially in line with previous studies of the labor market. We find that those with higher wages move less often (but see Lindeboom and Theeuwes 1991) and that commuting distance positively affects job mobility. These results correspond to results, *inter alia*, by Zax (1991), Van den Berg (1992), and Van Ophem (1991). Furthermore, the variables of age, size of the branch, number of subordinates, educational level, and calendar year are found to be statistically significant.¹³ These results do not need further discussion. Finally, we found that unobserved variables play a role as an explanation for the observed moving behavior. According to the results, 2 percent of the workers have about 4.5 times higher hazard rates of changing jobs than do the other workers ($v_1 = 1.37$, $v_2 = 6.99$).

The empirical model has been specified and interpreted in terms of hazard rates. Clearly, there is an inverse relationship between job hazard rates and job durations. Given the specification of the log likelihood function, it can be shown that $\log(T) = -X\beta + \varepsilon$, where T is the job duration, ε is a random variable, and X and β are the same as defined in equation 2 (see Lancaster 1990). So, the estimated β can be interpreted as a quasi-elasticity of the job duration with respect to the explanatory variables X . For example, the quasi-elasticity of the job duration with respect to commuting distance is 0.47. So, an increase of 10 kilometers in commuting distance reduces the expected duration by 37.5 percent (since $\exp^{(\log(T) - 0.47)}/T = \exp^{(-0.47)} = 0.625$).

THE ROBUSTNESS OF THE ESTIMATION RESULTS

Now we will examine whether the empirical results as presented above in the Estimates section are sensitive concerning the model specification chosen.

First, it has been assumed that the mixing distribution $h(v)$ is parameterized with two discrete mass points. Hence, we have reestimated the model assuming that the mixing distribution is parameterized with three discrete mass points. It appeared that the results were identical.

Second, we have assumed that the effect of the spatial location of the spouse's job can be measured by means of the variables z_2 and z_3 . However, as an alternative, one may use α , which represents the angle between z_1 and z_2 . Clearly, α and z_3 are positively, but nonlinearly, related to each other, conditional on z_1 and z_2 (by the law of cosines, $z_3^2 = z_1^2 + z_2^2 - 2 \cdot z_1 \cdot z_2 \cdot \cos(\alpha)$, $0 < \alpha < \pi$; $\partial z_3 / \partial \alpha = 2 \cdot z_1 \cdot z_2 \cdot \sin(\alpha) \geq 0$). We have therefore reestimated the model using α instead of z_3 (model 2). It appeared that the effect of α is positive. Nevertheless, the results show that the effect of α is not significant at the 5 percent level. Given this specification, the coefficient of the commuting distance of the spouse is insignificant at the 5 percent level.¹⁴ Hence, although both specifications provide evidence that the spatial location of the spouse's job affects job mobility, it is not clear which factor (viz. z_2 , z_3 , or α) is the main cause.¹⁵ In conclusion, we are able to provide some evidence against the null hypothesis that the spatial location of the spouse's job does not affect job mobility. Nevertheless, we fail to provide evidence of how the spatial location of the spouse's job affects job mobility. More decisive results may be expected with larger data sets.

Third, one may argue that single-wage earners and those in two-earner households behave in structurally different ways, which cannot be captured only by means of one single regressor, two-earner household. Therefore, the model is reestimated on two subsets of observations. We have reestimated the model given a subset of observations of workers who live with spouses (who may or may not be employed) and given a subset of observations of workers who live with employed spouses (models 3 and 4). In the latter estimations, we have restricted the range of explanatory variables, since the number of observations is limited. It appears that the estimation results do not contradict those presented above in the Estimates section (however, since many coefficients are insignificant, the power of this test is limited). The only difference is that the results of model 3 indicate that job mobility is reduced by the presence of a working spouse for females as well as for males. Nevertheless, the effects of the coefficient for the explanatory variable male in model 3 and model 4 indicate that females with spouses and females with employed spouses change jobs less often, so the original conclusion that females who belong to two-earner households change jobs less often is not invalidated. Hence, we conclude that the results presented are robust with the chosen specification.

CONCLUSION

We have tested the hypothesis that on-the-job moving behavior differs for those in two-earner households and single-wage earners. Given a data set of two-earner households and single-wage earners in the Netherlands, we found that female workers with spouses, particularly when they belong to two-earner households, tend to change jobs less often than do other workers. This might be interpreted as a sign that many female workers with spouses refuse job offers that are accepted by

other workers. The empirical results indicate that job mobility does not depend strongly on the spatial location of the spouse's job. It would be interesting to see whether studies outside the Netherlands confirm these results.

APPENDIX VARIANCE OF ESTIMATES

Given random variables X_1, \dots, X_n , the variance of the sum of the random variables can be written as $\left(\sum_1^n X_i\right) = \sum_1^n \text{var } X_i + 2 \sum_i \sum_{i < j} \text{cov}(X_i, X_j)$.

The covariance matrix of the estimates β for (1) a spouse, (2) two-earner, (3) two-earner (if female), and (4) spouse (if female) has been estimated as

$$\begin{bmatrix} 0.04 & & & \\ -0.02 & 0.08 & & \\ 0.02 & -0.07 & 0.14 & \\ -0.04 & 0.03 & -0.07 & 0.14 \end{bmatrix}.$$

The variance of the sum of an effect can then be calculated using the above formula and the estimated covariance matrix. For example, the variance of the presence of an employed spouse for female workers is equal to $0.04 + 0.08 + 0.14 + 0.14 + 2(-0.02 + 0.02 - 0.04 - 0.07 + 0.03 - 0.07)$. So the standard deviation is equal is 0.32.

NOTES

1. Data presented in Linneman and Graves (1983) for the United States indicate that only 11 percent of the heads of households who changed jobs also changed county of residence in the same year. The percentage of moves that may be interpreted as a migration is much less, since a change of county does not necessarily imply a move over a large distance: according to the same data, 50 percent of the heads of households who changed county of residence did not change jobs in the same year. In the Netherlands, the average distance between the old and the new residence is 4.2 kilometers, while 75 percent of all residential moves are less than 15 kilometers (van Dijk 1986).

2. We do not distinguish between voluntary job-to-job moves and voluntary job-to-unemployment moves.

3. A rationale of this definition of a two-earner household is that it is generally thought that the labor market behavior of a worker who does not work full-time does not influence the labor market behavior of the worker's spouse.

4. Both wage earners of a two-earner household are among the 122 observations. So, the job mobility behavior of these workers is dependent. Fortunately, this type of dependency does not affect the consistency of the estimates and is ignored in the current article.

5. We observe 3,020 annual spells generated by 589 individuals. The dependency between the spell observations is taken into account by the unobserved variable v .

6. When we allowed for duration dependence by means of a Weibull model, the model did not converge.

7. Characteristics that are only defined for two-earner households are set to zero in the case of single-wage earners.

8. This seems to be a plausible assumption for the average worker in the Netherlands. The median distance is about 20 kilometers. So, we assume that jobs are offered within a range of 100 kilometers.

9. Since the commuting distance is observed with a potentially large measurement error, the reported effect of commuting distance is an underestimate. Furthermore, the survey does not contain information on commuting time. This is unfortunate, since it has been shown that, in the Netherlands, commuting time has a stronger effect on job search behavior than commuting distance has (van Ommeren 1996). In addition, Dubin (1991) has shown for the United States that workers react stronger to changes in commuting time than to changes in commuting distance.

10. In this article, we assume that job-moving behavior does not differ for males and females except for a number of gender-dependent dummies (see Viscusi 1980). We have tested the hypothesis that job-moving behavior differs for males and females by estimating the models for males and females separately. The sum of the log likelihood of these two models is about 15 higher than those of the models presented in this paper. Using a standard likelihood ratio test, the hypothesis that job-moving behavior differs for males and females cannot be rejected.

11. In a previous version of this article, the presence of an employed spouse was assumed not to be gender dependent. This assumption was criticized by one of the referees. It seems indeed more accurate to make this explanatory variable gender dependent.

12. An alternative explanation would be that single-wage earners receive more job offers than those in two-earner households, for example, because single-wage earners might be more productive according to the employers. The latter explanation, however, is, as far as we know, not supported by empirical facts.

13. We have also estimated the model using annual dummies. Given this specification however, the covariance matrix, computed as the inverse of the computed Hessian, failed to invert.

14. In addition, in model 2, the coefficient of commuting distance z_1 is not significant at the 5 percent level. However, this coefficient seems to be less pronounced due to the collinearity of the regressors commuting distance and commuting distance of spouse. We have reestimated the model excluding the regressor commuting distance of spouse. The estimates were hardly affected by this exclusion, except for the coefficient of commuting distance that became significant at the 5 percent level.

15. To explore this result further, we have tested whether the specification that includes α or the specification that includes z_3 is more appropriate. Because the two specifications are nonnested, a statistical encompassing test of the specifications is used (Mizon and Richard 1986). It appears that at the 5 percent significance level, one cannot distinguish between the two specifications (both specifications are not rejected).

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